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Currency hedging and goods trade

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Abstract

A puzzle in empirical international finance is the difficulty in finding a large and negative effect of exchange rate volatility on international trade. A common explanation is the availability of hedging instruments. This paper examines the empirical validity of this explanation using data on over 1000 country pairs. Which countries have currency hedging instruments is not perfectly observable. This paper deals with the problem by specifying an endogenous regime-switching regression. There are two main findings. First, there is no evidence in the data to support the validity of the hedging hypothesis. Second, for country pairs with large trade potential, exchange rate volatility deters goods trade to an extent much larger than that typically has been documented in the literature (without using the switching regression specification). © 1999 Published by Elsevier Science B.V. All rights reserved.

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1. Introduction

A puzzle in empirical international finance is the difficulty in identifying a large and negative effect of exchange rate volatility on trade.[1] This has led to

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[1] There is a large number of empirical papers on the subject. Generally speaking, it is difficult to find a negative effect that is statistically significant and/or quantitatively large. For good survey papers, see Farrell et al. (1983), IMF (1984) and Belanger and Gutierrez (1990).

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a bifurcation of reactions. On the one hand, policy circles choose to ignore this literature, and continue to believe that exchange rate volatility has a large and negative effect on goods trade. For example, government officials in Europe explicitly and repeatedly cite this effect as a primary justification for the European Monetary System (which reduces exchange rate volatility) and for the drive for a single currency in Europe (which eliminates the volatility for member countries).

On the other hand, clever economists start to think of clever explanations for why the effect should be small on a conceptual level. Among the explanations, the availability of hedging instruments is often proposed as a solution to the puzzle.[2] For example, a recent survey paper by Cote (1994, p. 6) clearly states that 'the availability of forward cover reduces the effect of exchange rate volatility'.

In the vast empirical literature on the effect of exchange rate volatility on trade, a small subset (e.g., De Grauwe and de Bellefroid, 1986) did find a negative and statistically significant effect. They tend to be studies using data in the 1970s or early 1980s. Using both early and more recent data, Frankel and Wei (1994) found a negative coefficient before the mid-1980s, but the negative effect has disappeared since then. This pattern is in principle consistent with the hedging hypothesis: currency hedging products were not as well developed in the 1970s and early 1980s as they are now. In fact, currency options were first traded in the U.S. around 1982–1983 and have since grown exponentially in both volumes and varieties.

Although the hedging hypothesis seems intuitive and almost self-evident, to my knowledge, it has not been subjected to a careful examination. A primary objective of this paper is to undertake such an investigation. It may seem that such a test should be straightforward, as the hypothesis implies that country pairs with hedging instruments should experience a lower effect of volatility on trade than country pairs without them. However, it is not easy to separate the two groups. First, as many hedging instruments are over-the-counter products, they may not always be reported by national governments or international institutions. This leads to an under-counting of country pairs that can hedge. On the other hand, in certain countries where a hedging market is reportedly in existence, the restrictions can be so severe that exporters and importers do not find it of practical use. This may lead to an over-counting of country pairs that have access to hedging instruments. The two reasons imply that the numbers in

[2] There are other explanations, which this paper does not focus on. (a) Convexity in the profit function may be large enough to offset concavity in the utility function of firm owners, so that firms do not care about exchange rate volatility (Caballero and Corbo, 1989). (b) The extent of exchange rate volatility (for five industrialized countries) is not large enough for modestly risk averse producers to care about its effect (Gagnon, 1993). (c) The exchange rate volatility may in fact offset other types of business risks (Makin, 1978).

official reports may not be reliable. With this empirical difficulty in mind, part of the innovation of this paper is to design a statistical rule that endogenously separates the two groups based on observable characteristics. To reveal the punch line up front, the evidence that I will present denies the validity of the hedging hypothesis.

I will organize the paper in the following way. Section 2 provides a discussion of the hedging hypothesis, and some motivation for the empirical specification used in a later section. Section 3 describes the data. The meat of the paper is in Section 4, which provides empirical tests of the hedging hypothesis, using data on over 1000 pairs of bilateral trading partners during 1975–1990. Finally, Section 5 offers some concluding thoughts.

2. Discussion of the hedging hypothesis

If hedging instruments on exchange rate (e.g., forwards) are costlessly available, then firms' production and exports are not affected by exchange rate volatility. This was first proved by Ethier (1973) and extended by a number of authors including Kawai and Zilcha (1986). This theoretical result, sometimes known as the separation theorem, is the logical foundation of the hedging hypothesis.

Over the last twenty years, hedging instruments for exchange rate risks have rapidly proliferated. Aside from currency futures that are traded on organized exchanges, there are bank-offered forward contracts and currency swaps. Since 1978, currency options have been quickly developed into a high-volume liquid market.[3] This observation together with the separation theorem lends plausibility to the hypothesis that the increasing availability of hedging instrument is responsible for the diminished effect of exchange rate volatility on trade.

We should note that the hedging hypothesis as stated needs to be qualified in a number of dimensions. First, and perhaps most importantly, the use of hedging instruments is not costless. Not only the cost is not constant, it is often positively related to the exchange rate volatility (e.g., through changes in the premium on currency options, or the bid–ask spread on forward contracts). Second, an increase in volatility may affect (often depress) trade indirectly through its effect on the forward rate (Viaene and de Vries, 1992). Third, hedging instruments are often available only for short horizons (typically one month to

[3] Currency options were first introduced in 1978 on the European Options Exchange. The North American trading began in late 1982 on the Philadelphia Exchange and expanded in 1985 to include Chicago Mercantile Exchange. In addition to the products mentioned in the text, there are also variations of the basic instruments or 'exotic' products. They include cylinders, collars, zero premium options, G-hedges, compound options, break forwards, participating forwards, and extra, scout, pooled and Asian options.

a year), which may be shorter than the planning horizon of many exporters and importers. Fourth, exporters and importers may care about real exchange rate risk, whereas available hedging instruments are designed to hedge against nominal exchange rate risk. These qualifications undermine the explanatory power of the hedging hypothesis for the observation of a small effect of volatility on trade. For the moment, let us give the maximum benefit of doubt to the hedging hypothesis by assuming away these qualifications.

Many countries with flexible exchange rates do not have well-developed currency hedging market. This observation is going to be important in our test of the hedging hypothesis in the next section. To make use of the observation, we have to have an idea about how hedging instruments come to life. In discussing financial innovation in general, and the emergence of futures market in particular, both Miller (1986) and Telser (1991)[4] emphasized potential liquidity as the crucial determinant for the success of a financial instrument. That is, only a derivative product with enough potential demand for it and hence liquidity will survive in the market. Otherwise, low liquidity translates into a high trading cost. Historically, many futures or options instruments were invented, or even started trading in an organized exchange, but eventually died out due to low liquidity.

This theory of when a hedging instrument would emerge and sustain in the market suggests a non-linear relationship between currency volatility and goods trade in a cross-section context. Those country pairs which have small actual or potential trade would not develop a liquid market for currency hedging instruments. Even though hedging products can be developed by banks, or synthetically manufactured with other financial products by firms, they would be too costly to be widely used in practice. Consequently, these countries' trade may be depressed by the exchange rate volatility. On the other hand, for country pairs with a large trade potential, currency hedging instruments will emerge and be available at a low cost, the exchange rate volatility may no longer depress goods trade.[5]

To my knowledge, no previous empirical studies on the effect of exchange rate volatility on trade has investigated this non-linearity. Under the hedging hypothesis, all cross-sectional studies that do not take into account the non-linearity potentially suffer from a misspecification and bias towards zero the estimated effect of exchange rate volatility on trade.

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[4] See also Fry (1994).

[5] The liquidity theory of hedging instrument development permits, in principle, multiple equilibria. It is not discussed in either Miller (1986), Telser (1991) or Fry (1994). The multiple equilibria possibility does not arise for large trading countries or small trading countries, but only for 'intermediate' cases. In this paper, I will ignore this possibility for simplicity.

3. Data

The basic dependent variable is bilateral trade among 63 countries in the world for the years 1975, 1980, 1985 and 1990. The data on the first three years are from the United Nations' Trade Matrix, and that on 1990 is from the IMF's Direction of Trade Statistics. In principle, the 63 countries should give us 1953 pairs of bilateral trade. Due to missing data, I actually have between 1101 and 1453 observations for each of the four years, and 5542 observations for panel regressions.

The following data are extracted from the IMF's International Financial Statistics data base: nominal exchange rates (monthly, end of period, line ae), CPI index (line 64), money supply (M2, line 34 plus line 35), GNP (line 99a) or GDP (line 99b), and population (line 99z).

The measure of real exchange rate volatility in this paper is the standard deviation of the first difference in the log of the monthly exchange rates over a two-year period (current and preceding years). Later in the paper, I will experiment with an instrumental variables approach as well as nominal exchange rate volatility.

Distance between countries are computed as their 'greater circle distance'. *Adjacency* is a dummy for country pairs sharing a common land border. *Common Link* is a dummy for country pairs that share a common language (any of the nine major languages: Arabic, Chinese, Dutch, English, French, German, Japanese, Portugese, and Spanish) and/or colonial tie. They are available at Shang-Jin Wei's web page: www.nber.org/~wei.

4. Empirical examination

In this section, we turn to an empirical examination of the following joint hypothesis: (1) The volatility elasticity of trade is negative and the same for all country pairs in the absence of hedging instruments; (2) the elasticity is reduced (possibly to zero) for country pairs that have access to hedging instruments.

4.1. First look: OLS regressions based on reported hedging instruments

I start the investigation with a naive approach. Let me construct a dummy variable, FORWARD, for country pairs for which hedging instruments on their bilateral exchange rates are reportedly available according to two sources, International Monetary Fund (IMF) and Data Resource Inc. (DRI). The futures and options are traded on organized exchanges. The *International Capital Market* (IMF, 1993, August) lists ten major organized exchanges that trade these instruments. They are located in the United States (4), Japan (1), Spain (1), Netherlands (2), Brazil (1) and Singapore (1). Collectively, they cover 13 bilateral exchange rates (including the ECU), mostly against the U.S. dollar. Currency forwards and swaps are between banks or between banks and corporate customers. A comprehensive list is hard to compile. The database from the Data Resource Inc. reports forward data on 26 currencies against the U.S. dollar.

In the regression, I specify

$$y_{ij} = X'_{ij}\beta + \Theta z_{ij} + \delta z_{ij}\,\mathrm{FORWARD}_{ij} + u_{ij}$$

where y is total trade between countries i and j in logarithm (the country pair subscript ij in this and following sentences is omitted for convenience). X is a vector of variables (other than exchange rate volatility) that determines the volume of trade. z is the exchange rate volatility. u is a random variable that is i.i.d. normal with mean zero and variance σ_u.

If the hedging hypothesis is correct as an explanation for the low volatility effect on trade, one should observe the following values for the two key parameters: $\Theta < 0$ and $\delta = -\Theta > 0$. That is, for country pairs that do not have hedging instruments, volatility has a depressing effect on trade. For country pairs that can hedge, the effect disappears.

There are several ways to specify the X vector. In the actual implementation, I use a gravity specification and include six terms: a constant, the product of the two countries' GNPs, the product of the two countries' per capita GNPs, the distance between the (economic centers of the) two countries, a dummy variable for countries that share a common land border, and finally, a dummy variable for country pairs that share a common language or some historic/colonial tie. All variables except the dummies and the constant are in logarithm. The choice of the variables and the specification are taken from Frankel and Wei (1994).

The gravity specification can be justified by a differentiated product framework with increasing returns to scale (Helpman and Krugman, 1985) or by the neo-classic Ricardian or Heckscher–Ohlin theories (Deardorff, 1995). The gravity regressions tend to perform remarkably well empirically (Deardorff, 1984). A typical regression in those studies can produce an adjusted R^2 in excess of 70%, a remarkably large explanatory power for cross-section data. This is much better than a typical regression that has only factor endowment terms. Frankel and Wei (1994) report that adding factor endowment terms to a gravity specification does not noticeably improve the adjusted R^2.

For the purpose of comparison, let me first report some regression results without the FORWARD dummy (Table 1). The first four columns are year-by-year regressions. We notice that the effects of exchange rate volatility were negative in 1975 and 1980 (-1.72 and -7.51, respectively), and were statistically significant. It became indifferent from zero in 1985 and even positive in 1990. As stated in the beginning of the paper, this pattern would be consistent with the hedging hypothesis since the trading volume on currency derivatives has increased steadily and substantially over the last two decades. The last column reports a fixed-effects regression where data from all four years are pooled. In

Table 1
Currency volatility and goods trade: OLS estimates

	1975	1980	1985	1990	Panel
Volatility	− 1.72*	− 7.51*	0.55	1.85*	1.69*
θ	(0.63)	(1.21)	(0.47)	(0.44)	(0.32)
GNP	0.71*	0.75*	0.53*	0.79*	0.71*
β_1	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)
GNP/capita	0.24*	0.23*	0.07*	0.11*	0.18*
β_2	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)
Distance	− 0.61*	− 0.64*	− 0.43*	− 0.57*	− 0.58*
β_3	(0.04)	(0.03)	(0.04)	(0.03)	(0.02)
Adj	0.65*	0.68*	0.62*	0.89*	0.72*
β_4	(0.16)	(0.15)	(0.21)	(0.20)	(0.08)
Common links	0.46*	0.61*	0.53*	0.67*	0.58*
β_5	(0.09)	(0.09)	(0.10)	(0.09)	(0.05)
Constant	− 9.14*	− 9.94*	− 6.33*	2.40*	N.R.
β_0	(0.54)	(0.52)	(0.59)	(0.34)	
# obs	1230	1457	1101	1307	5542
Log likelihood	− 1822.86	− 2208.21	− 1754.46	− 1903.1	− 8830.2
R^2	0.73	0.73	0.48	0.76	0.74

Note: (1) The panel has year and region (Asia, Western Hemisphere, Europe, Africa) dummies. (2) N.R. = Not reported.

this regression, the effect of the exchange rate volatility is positive (1.69) and statistically significant.

Now let me turn to the regression results (Table 2) with an added interactive term between the FORWARD dummy and the exchange rate volatility. If the hedging hypothesis is correct *and* the IMF and DRI accounts of which currency has hedging instruments is accurate, then the coefficient on the interactive term, δ, should be positive and statistically significant. What does the data tell us? In the two years that the volatility does have a negative effect on trade (1975 and 1980), there is no statistically significant difference between the country pairs that have hedging instruments and those that do not. In fact, contrary to the null hypothesis, the point estimates for the coefficient on the interactive dummy, δ, are negative for the two years. In the only year (1985) that δ is positive and significant, there is no discernible depressing effect of volatility on trade even for country pairs that do not have hedging instruments. Hence, the evidence does not support the null hypothesis that the availability of the hedging instrument is responsible for the diminished effect of volatility on trade. In the fixed-effects panel regression (last column in Table 2), the coefficient on the interactive term, δ, is indeed positive (7.03) and statistically significant at the five percent level. On the other hand, the effect of exchange rate volatility on trade is not negative in this specification even for country pairs without the hedging instruments.

Table 2
Elasticity of trade with respect to exchange rate volatility: The role of hedging instruments (1975–1990)

	1975	1980	1985	1990	Panel
Dependent variable: ln(trade)					
Volatility	− 1.79*	− 7.66*	0.29	3.12*	1.72*
	(0.72)	(1.23)	(0.55)	(0.35)	(0.32)
FORWARD volatility	− 0.48	− 0.08	19.67*	− 0.19*	7.03*
	(5.83)	(6.25)	(4.87)	(5.75)	(3.04)
GNP	0.74*	0.78*	0.57*	0.80*	0.71*
	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)
GNP/capita	0.25*	0.25*	0.07*	0.11*	0.19*
	(0.02)	(0.02)	(0.03)	(0.02)	(0.01)
Distance	− 0.69*	− 0.66*	− 0.50*	− 0.63*	− 0.58*
	(0.04)	(0.04)	(0.05)	(0.04)	(0.02)
Adjacency	0.59*	0.71*	0.70*	0.93*	0.73*
	(0.16)	(0.16)	(0.20)	(0.17)	(0.08)
Common	0.46*	0.63*	0.63*	0.60*	0.59*
Links	(0.10)	(0.09)	(0.10)	(0.09)	(0.05)
Year dummies					Yes
Region dummies					Yes
# obs.	1230	1457	1101	1307	5542
adj. R^2	0.73	0.73	0.49	0.76	0.74
Log likelihood	− 2043.1	− 2362.7	− 2101.3	− 2253.5	− 8793.3

Note: (1) Heteroskedasticity-robust standard errors are reported below the coefficient estimates. (2) All regressions have an intercept which is not reported here. All variables except the dummies and exchange rate volatility are in natural logarithm. (3) Ex-volatility is the real exchange rate volatility, measured as the standard deviation of the monthly real exchange rates over a two-year (current and last years) period. (4) FORWARD is a dummy for country pairs that, according to the Data Resource Inc., have forward contracts (or other hedging instruments) available directly on their bilateral exchange rate.
* Statistically significantly different from zero at 5% level.

One may worry that our results may be spurious because countries without hedging instruments could have substantially smaller exchange rate variability. In the extreme, if all countries that do not having hedging instruments also peg their exchange rate to one of the major currencies, their exchange rate volatility (zero by definition) would have no correlation with their trade. To check this possibility, I went through the exchange rate arrangement list of the *International Financial Statistics* published by the International Monetary Fund, and found that very few countries in our sample proclaim a fixed exchange rate system.

Table 3 reports the average exchange rate volatility for the whole sample, and for the subsamples grouped according to whether or not the country pairs have hedging instruments. For all four years, the average real exchange rate volatility

Table 3
Summary statistics: Exchange rate volatility, 1975–1990

	1975	1980	1985	1990
Real exchange rate volatility				
Whole sample				
Mean	0.0418	0.0351	0.0510	0.0611
Std dev.	(0.0397)	(0.0265)	(0.0670)	(0.0820)
# obs	1316	1503	1241	1494
Country pairs with hedging instruments as reported by DRI & IMF				
Mean	0.0274	0.0253	0.0324	0.0229
Std dev.	(0.0076)	(0.0087)	(0.0141)	(0.0093)
# obs	21	23	24	24
Country pairs without hedging instruments				
Mean	0.0421	0.0352	0.0513	0.0617
Std dev.	(0.0399)	(0.0267)	(0.0676)	(0.0825)
# obs	1295	1480	1217	1470
Nominal exchange rate volatility				
Whole sample				
Mean	0.0373	0.0303	0.0514	0.0738
Std dev.	(0.0448)	(0.0274)	(0.0735)	(0.1112)
# obs	1316	1503	1241	1494
Country pairs with hedging instruments				
Mean	0.0253	0.0241	0.0321	0.0217
Std dev.	(0.0093)	(0.0093)	(0.0143)	(0.0105)
# obs	21	23	24	24
Country pairs without hedging instruments				
Mean	0.0375	0.0304	0.0517	0.0746
Std dev.	(0.0451)	(0.0275)	(0.0741)	(0.1119)
# obs	1295	1480	1217	1470

for country pairs that do not have hedging instruments is actually *higher* than for those pairs that do. In 1975, for example, the average real exchange rate volatility was 0.0418 for the whole sample. The volatility for country pairs that have hedging instruments was only 0.0274, compared with 0.0421 for country pairs without hedging instruments. The same pattern holds for all other years and for nominal exchange rate volatility.

4.2. Endogenous switching regression

The above approach may be too naive. The IMF and DRI account of which countries have hedging instruments and which do not may not be accurate for the two reasons stated in the first section of the paper. In this subsection, I will specify an equation that infers which country pairs are likely to have developed hedging instruments based on potential liquidity for such a market. Let us hypothesize that the data is in fact from two regimes. Regime I (denoted by $R = 1$) includes all countries that do not have any access to a hedging instrument. Regime 2 (denoted by $R = 2$) includes all country pairs that do have some hedging instruments.

$$y_{ij} = X'_{ij}\beta + \theta_{ij}z_{ij} + e_{ij}$$

where

$$\theta_{ij} = \begin{cases} \theta_1 & \text{if } R = 1, \\ \theta_2 & \text{if } R = 2. \end{cases}$$

Define y^* as the virtual trade, or the level of trade (in logarithm) in the absence of any hedging instrument, plus a random variable with zero mean. That is,

$$y^*_{ij} = X'_{ij}\beta + \Theta_1 z_{ij} + u_{ij}$$

where u_{ij} is an i.i.d. random variable with mean zero and variance σ_u. The correlation coefficient between e and u is ρ. Let us assume that a hedging market will emerge if and only if y^* exceeds a critical value, Γ. That is,

$$R_{ij} = \begin{cases} 1 & \text{if } y^*_{ij} < \Gamma, \\ 2 & \text{if } y^*_{ij} \geq \Gamma. \end{cases}$$

I do not force u_j to be equal to e_j, because the emergence of a hedging market may also be affected by other random noises (represented here by $u_{ij} - e_{ij}$) unrelated to the level of trade.

Let me use ϕ and Φ to denote the probability density function and the cumulative distribution function of a standard normal random variable, respectively.

$$\begin{aligned} \text{Prob}\,(R = 1) &= \text{Prob}\,(y^* < \Gamma) = \text{Prob}\,(e < \Gamma - X'\beta - \Theta_1 z) \\ &= \Phi((\Gamma - X'\beta - \Theta_1 z)/\sigma_u), \\ \text{Prob}\,(R = 2) &= \text{Prob}\,(y^* \geq \Gamma) = 1 - \text{Prob}\,(R = 1). \end{aligned}$$

The likelihood of observing observation 'ij', L_j, is given by

$$L_{ij} = f(e_{ij} \mid R = 1)\, \text{Prob}\,(R = 1) + f(e_{ij} \mid R = 2)\, \text{Prob}\,(R = 2).$$

With some algebra (see the appendix for details), one can show the following:

$$f(e_{ij} \mid R = 1) = \frac{1}{\sigma_e}\,\phi\left(\frac{e_{1ij}}{\sigma_e}\right)\Phi\left(\frac{u/\sigma_u - \rho\,(e_{1ij}/\sigma_e)}{\sqrt{1-\rho^2}}\right)\Phi^{-1}\left(\frac{c_{ij}}{\sigma_u}\right)$$

and

$$f(e_{ij} \mid R=2) = \frac{1}{\sigma_e}\,\phi\left(\frac{e_{2ij}}{\sigma_e}\right)\left[1-\Phi\left(\frac{u/\sigma_u - \rho\,(e_{2ij}/\sigma_e)}{\sqrt{1-\rho^2}}\right)\right] \times \left[1-\Phi\left(\frac{c_{ij}}{\sigma_u}\right)\right]^{-1},$$

where

$$c_{ij} \equiv \Gamma - X'_{ij}\beta - \Theta_1 z_{ij}, \quad e_{1ij} \equiv y_{ij} - X'_{ij}\beta - \Theta_1 z_{ij},$$

$$e_{2ij} \equiv y_{ij} - X'_{ij}\beta - \Theta_2 z_{ij}.$$

The log likelihood function, L, is

$$L = \Sigma_{ij} \log (l_{ij}).$$

The parameters, β, Θ_1, Θ_2, Γ, σ_u, σ_e and ρ can be estimated by maximizing the log likelihood function. Notice that in usual switching regressions, the parameter σ_u and those contained in c_j are not jointly identified. Here, because there are strong parameter restrictions (the same β vector in both the main and the switching regressions), all 12 parameters are identified. On the other hand, under the hypothesis that $\Theta_1 = \Theta_2$, the parameters Γ, σ_u, and ρ are not identifiable (but Θ and other parameters can still be estimated).

Table 4 reports the results of the estimation of the endogenous regime switching regression. The results are striking. For country pairs that have small virtual trade (trade in the absence of any hedging instrument), the volatility depresses trade only in 1975 and 1980; but the negative effect disappears in later years. In sharp contrast, for country pairs that have large virtual trade, exchange rate volatility seems to have numerically large and statistically significant negative effect on trade throughout the sample. In 1990, for example, a 1% increase (i.e., by 0.01) in the exchange rate volatility was associated with a 10% drop in the corresponding bilateral trade.[6] When I perform a Wald test on the hypothesis that $\theta_1 - \theta_2 = 0$, I can reject the hypothesis for all years at the 5% level. [For example, the chi-square statistics are 93.18 and 9.23 for 1980 and 1985, respectively, exceeding the critical value at the 5% level.][7] In Column 5, a fixed-effects regression is estimated.[8] Here again, the estimate on the volatility

[6] Using data on ten industrialized countries, De Grauwe (1988) found that real exchange rate volatility has a significantly negative effect on the growth rate of trade.

[7] Note that under the null hypothesis of $\theta_1 = \theta_2$, Γ, σ_u, and ρ are not identifiable (any values are equally likely as any other values). But the estimator of θ has a well-defined distribution (and the likelihood function still has a finite maximum). Therefore, we can construct a (Wald or likelihood ratio) test for the hypothesis $\theta_1 = \theta_2$.

[8] The fixed-effects specification includes year and region (Asia, Western Hemisphere, Europe and Africa) dummies whose estimates are reported. We do not include all country-pair fixed effects because the resulting large number of extra parameters (1953 in total) causes problems in convergence in the estimation.

Table 4
Endogenous switching regressions

	1975	1980	1985	1990	Panel I	Panel II	Panel III
Volatility 1	− 1.32*	− 5.59*	0.57	2.41*	2.37*	2.27*	2.24*
θ_1	(0.48)	(1.28)	(0.62)	(0.48)	(0.30)	(0.33)	(0.34)
Volatility 2	− 2.29*	− 13.66*	− 29.45*	− 10.31*	− 33.87*	− 34.27*	− 33.64*
θ_2	(0.58)	(2.45)	(9.78)	(2.74)	(2.31)	(2.38)	(3.83)
Volatility 3						− 2.59	− 3.98
θ_3						(3.34)	(3.94)
GNP	0.72*	0.77*	0.53*	0.81*	0.74*	0.74*	0.76*
β_1	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
GNP/capita	0.24*	0.23*	0.07*	0.10*	0.18*	0.18*	0.19*
β_2	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
Distance	− 0.62*	− 0.64*	− 0.42*	− 0.60*	− 0.59*	− 0.59*	− 0.55*
β_3	(0.04)	(0.04)	(0.05)	(0.04)	(0.02)	(0.02)	(0.03)
Adjacency	0.65*	0.73*	0.64*	0.89*	0.75*	0.74*	0.72*
β_4	(0.17)	(0.16)	(0.19)	(0.18)	(0.09)	(0.09)	(0.09)
Common links	0.46*	0.63*	0.53*	0.70*	0.62*	0.62*	0.62*
β_5	(0.09)	(0.08)	(0.10)	(0.08)	(0.05)	(0.05)	(0.05)
Trade blocs							0.35*
β_6							(0.08)
Critical value γ	4.65*	8.80*	15.03*	9.15*	10.30*	10.20*	10.14*
	(0.07)	(1.15)	(6.19)	(0.81)	(0.42)	(0.44)	(0.47)
σ_u	1.06*	4.68*	5.18	3.09*	3.78*	3.87	3.88*
	(0.03)	(1.21)	(3.20)	(0.71)	(0.29)	(0.33)	(0.34)
σ_e	1.07*	1.16*	1.22*	1.10*	1.26*	1.26*	1.25*
	(0.02)	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)
ρ	0.99*	0.93*	0.96*	0.86*	0.87*	0.87*	0.86*
	(0.01)	(0.04)	(0.05)	(0.06)	(0.02)	(0.02)	(0.02)
Constant	Yes	Yes	Yes	Yes			
Year dummies					Yes	Yes	Yes
Region dummies					Yes	Yes	Yes
# obs	1230	1457	1101	1307	5542	5542	5542
Log likelihood	− 1820.8	− 2197.8	− 1749.0	− 1890.3	− 8746.4	− 8746.0	− 8703.4

Note: (1) Standard errors are reported in parentheses.
* Statistically significantly different from zero at 5% level by a t-test.

effect for country pairs with large virtual trade is a negative number that is statistically significant and quantitatively large. These estimates strongly contradict the hypothesis that attributes the abatement of a negative effect of exchange rate volatility to the availability of hedging instruments.

In the previous specification, we let a country pair develop hedging instruments only when its potential trade, y^*, exceeds a threshold. A possible extension is to allow the size of exchange rate volatility to play a role in the emergence of hedging instruments. For example, we may define a country pair as in the hedging regime ($R = 2$), if a combination of potential trade and the size of exchange rate volatility exceeds a threshold value, i.e., $(X'_{ij}\beta + \Theta_1 z_{ij} + u_{ij}) + \eta z_{ij} \geq \Gamma$. Otherwise, the country pair is in the no-hedging regime ($R = 1$). In this specification, if η is positive, then, hedging instruments are allowed to emerge even if the potential trade for a particular country pair is small as long as their exchange rate volatility is sufficiently large.

In actual implementation, this new specification is equivalent to define a country pair to be in the hedging regime ($R = 2$) if and only if $X'_{ij}\beta + \Theta_3 z_{ij} + u_{ij} \geq \Gamma$, where $\Theta_3 = \Theta_1 + \eta$. The estimation result based on this specification is reported as Column 6 (labelled as Panel II) in Table 4. As can be seen, the estimate of Θ_3 is not statistically different from zero. More importantly for our purpose, the estimates of Θ_1 and Θ_2 are virtually unaffected. In particular, contrary to the hedging hypothesis, for country pairs most likely to develop currency hedging products, the exchange rate volatility has a negative effect on trade.

In the empirical literature on bilateral trade patterns, there are studies that emphasize the role of regional trade blocs (e.g., Frankel and Wei, 1994). As an extension, I construct a dummy, 'Trade Blocs', for country pairs in a common preferential trade agreement (PTA), free trade area (FTA), customs union (CU), or common market (CM).[9] The last column of Table 4 reports the regression result with 'Trade Bloc' as an additional regressor. The dummy has a positive and significant coefficient, meaning that membership in a common trade bloc is associated with higher bilateral trade than otherwise. The coefficients on the volatility variables are similar to regressions with the dummy. In other words, the hedging hypothesis is still rejected by the data.

[9] In the sample, there are 217 country pairs that are in a common PTA or FTA, and 72 country pairs that are in a common CU or CM. These pairs may overlap with each other. The PTAs or FTAs in the sample are: ASEAN, Bangkok Agreement, EFTA, EC&EFTA, LAIA, NAFTA, Australia–New Zealand, Israel–EC, Israel–US, and Chile–Mexico. The CUs and CMs are: EC, Arab Maghreb Union, Andean Pact, MERCOSUR, and Arab Common Market. For detailed definitions of these trade blocs, see Fieleke (1992). In the construction of the dummy, if a particular country pair belongs to multiple trade blocs, the dummy still takes the value of one. Also, we have examined membership in formal agreements here. Not all of them are effective. Frankel and Wei (1994) also discussed possibly implicit trade blocs.

It may be useful to compare the estimated hedging opportunies from the endogenous switching approach with the FORWARD dummy constructed from the IMF/DRI information. Note first that the endogenous switching regression does not literally classify all country pairs into those with hedging opportunities versus those without. Rather, for every country pair, it estimates the probability of the pair in the hedging regime.

As an illustration, Table 5 reports the estimated probabilities of being in the hedging regime (ProbHedge) for all country pairs involving the United States on the one side. This is only a subset of all country pairs in the sample. Reporting all the results (over 1400 country pairs in total) would take up extra 20 pages of space without adding new insight. Furthermore, existing hedging instruments identified by the IMF/DRI information are mostly on dollar exchange rates.

There are several noteworthy features in Table 5. First, even though the endogenous switching specification does not directly use the FORWARD dummy (the IMF/DRI information), the latter contains predictive power for the former. In fact, if we regress the estimated hedging probability (ProbHedge) on a constant and the FORWARD dummy, the slope coefficient is positive and statistically significant at 1% level.

$$\text{ProbHedge} = \underset{(0.03)}{0.39} + \underset{(0.04)}{0.18}\ \text{FORWARD dummy}$$

$$R^2/\text{adj.}R^2 = 0.24/0.22, \quad \#\ \text{observations} = 58$$

where standard errors are in parentheses. The adjusted R-squared is 0.22, which is reasonable for a small cross-section regression like this. Second, the correspondence is far from perfect. In fact, the correlation coefficient between the two is 0.49. On the other hand, it is useful to stress again that the published IMF/DRI information on hedging instruments is problematic (which is why we use the endogenous switching regression to estimate the hedging probability in the first place).

4.3. Indirect hedging

The regime-switching specification so far focuses on only the possibility of direct hedging. An important extension is to allow cross-hedging through a common currency. For example, there may not be a direct hedging instrument between Taiwan and Brazil on their bilateral exchange rate, but they could nevertheless hedge against the exchange rate risk if each country can hedge risk on the bilateral rates between the U.S. dollar and their own currencies.

We will attempt to specify a more general model taking into account possible cross-hedging through the U.S. dollar. We will ignore cross-hedging through other major currencies for two reasons. First, this keeps manageable the algebra in deriving the likelihood function. Indeed, the allowance of cross-hedging

Table 5
Published information on versus estimated probability of hedging instruments (US as one side of the country pairs)

Country pair	IMF/DRI dummy	Estimated probability of hedging opportunities	Country pair	IMF/DRI index	Estimated probability of hedging opportunities
	FORWARD	ProbHedge		FORWARD	ProbHedge
Canada	1	0.819	Ecuador	0	0.245
France	1	0.724	Mexico	1	0.601
Germany	1	0.772	Peru	0	0.284
Italy	1	0.668	Venezuela	0	0.413
Japan	1	0.779	Bolivia	0	0.388
UK	1	0.782	Paraguay	0	0.234
Austria	1	0.592	Uruguay	0	0.278
Belgium	1	0.611	Algeria	0	0.402
Denmark	1	0.566	Nigeria	0	0.548
Finland	1	0.449	Egypt	0	0.359
Netherlands	1	0.638	Morocco	0	0.263
Norway	1	0.468	Tunisia	0	0.225
Sweden	1	0.552	Sudan	0	0.436
Switzerland	1	0.661	Ghana	0	0.476
Australia	1	0.527	Kenya	0	0.219
Greece	1	0.393	Ethiopia	0	0.168
Iceland	0	0.241	Iran	0	0.424
Ireland	1	0.493	Kuwait	0	0.360
New Zealand	1	0.387	Indonesia	0	0.277
Portugal	0	0.372	Hong Kong	1	0.381
Spain	0	0.577	India	0	0.506
South Africa	0	0.415	Korea	0	0.497
Turkey	0	0.971	Malaysia	1	0.340
Yugoslovia	0	0.612	Pakistan	0	0.320
Israel	0	0.417	Philippines	0	0.297
Argentina	0	0.398	Singapore	1	0.307
Brazil	0	0.719	Thailand	0	0.246
Chile	0	0.330	Hungary	0	0.432
Colombia	0	0.327	China	0	0.453

Note: The probability of hedging instruments is estimated based on an endogenous switching regression using 1980 data. The regression produces estimates for 1457 country pairs. This table only reports those estimates for the subset of country pairs involving the U.S. as one of the countries.

through one currency has already induced very complicated algebra. Second, almost all cross-hedging is indeed carried out through the U.S. dollar.[10]

We are going to classify all observations into five possible cases. Let R denote different cases. It takes one of the five values, 111, 121, 112, 122 and 2. '111'

[10] Conversations with currency traders have confirmed this impression.

represents the case that the virtual trade between country i and j, between i and the U.S. and between j and the U.S. are below the threshold. '121' and '112' represent the cases that the direct virtual trade between i and j, and one country's virtual trade with the U.S. are below the threshold, but the other country's virtual trade with the U.S. exceeds the threshold. '122' is the case in which both countries' virtual trade with the U.S. exceeds the threshold even though the virtual trade between them directly is too low relative to the threshold. And finally, '2' represents the case in which the direct virtual trade between i and j is above the threshold regardless of the size of their individual trade with the U.S.

The first three cases ($R = 111, 121$, or 112) correspond to the regimes in which no hedging instruments are available. The last two cases give rise to the regime in which a hedging market has been developed, either directly (when $R = 2$) or indirectly via the U.S. dollar (when $R = 122$). The (partial) elasticity of trade with respect to exchange rate volatility, θ, takes one of the two values, depending on whether or not a hedging market exists. To be precise, suppose the main regression is

$$y_{ij} = X_{ij}\beta + \theta_{ij} z_{ij} + e_{ij}.$$

Notice that the subscript, ij, is a *non-ordered* index for country pairs. By 'non-ordered', I mean that 'ij' and 'ji' are the same observation.

If the United States is on one side of a given country pair, ij, ($i = \text{US}$ or $j = \text{US}$), then,

$$\theta_{ij} = \begin{cases} \theta_1 & \text{if } R_{ij} = 1, \\ \theta_2 & \text{if } R_{ij} = 2, \end{cases}$$

and

$$R_{ij} = \begin{cases} 1 & \text{if } y^*_{ij} < \Gamma, \\ 2 & \text{if } y^*_{ij} \geq \Gamma. \end{cases}$$

The likelihood for a single observation ij is given by

$$L_{ij} = f(e_{ij} \mid R = 1) \Pr(R = 1) + f(e_{ij} \mid R = 2) \Pr(R = 2).$$

For all other country pairs ($i \neq \text{US}$ and $j \neq \text{US}$),

$$\theta_{ij} = \begin{cases} \theta_1 & \text{if } R = 111, 121 \text{ or } 112, \\ \theta_2 & \text{if } R = 2 \text{ or } 122, \end{cases}$$

$$R_{ij} = \begin{cases} 111 & \text{if } y^*_{ij} < \Gamma,\ y^*_{is} < \Gamma,\ y^*_{js} < \Gamma, \\ 121 & \text{if } y^*_{ij} < \Gamma,\ y^*_{is} \geq \Gamma,\ y^*_{js} < \Gamma, \\ 112 & \text{if } y^*_{ij} < \Gamma,\ y^*_{is} < \Gamma,\ y^*_{js} \geq \Gamma, \end{cases}$$

and

$$R_{ij} = \begin{cases} 122 & \text{if } y^*_{ij} < \Gamma,\ y^*_{is} \geq \Gamma,\ y^*_{js} \geq \Gamma, \\ 2 & \text{if } y^*_{ij} \geq \Gamma. \end{cases}$$

The likelihood for a single observation *ij* is given by

$$\begin{aligned} L_{ij} = & f(e_{ij} \mid R = 111)\Pr(R = 111) + f(e_{ij} \mid R = 121)\Pr(R = 121) \\ & + f(e_{ij} \mid R = 112)\Pr(R = 112) + f(e_{ij} \mid R = 122)\Pr(R = 122) \\ & + f(e_{ij} \mid R = 2)\Pr(R = 2). \end{aligned}$$

The probabilities of each of the five regimes are similar to the earlier case and can be worked out. The conditional densities are somewhat complicated and have no direct analog in the traditional switching regression literature. To minimize ugly-looking formulae in the text, I relegate all derivations of these terms to an appendix.

As before, our central interest is to test the following joint hypothesis: (1) the volatility elasticity of trade is the same (and negative) for all country pairs in the absence of any direct or indirect hedging markets; and (2) the existence of a hedging market eliminates the negative effect of volatility. In short, the null hypothesis is H_0: $\theta_1 < 0$ and $\theta_2 = 0$.

We again estimate the model with the maximum likelihood method. The results are reported in Table 6. Our central interest concerns the estimates of the two volatility elasticities of trade. The elasticities for country pairs with a *small* potential trade are not statistically different from zero for 1975, 1980 and 1985, and even positive and significant for 1990. On the other hand, the elasticities for country pairs with a *large* potential trade are negative and statistically significant throughout the sample. The panel regressions (the last three columns in the table) produce the same pattern. Once again, these estimates strongly reject the joint hypothesis.

4.4. Further robustness checks

A potential problem in the regressions reported so far is that the regressor exchange rate volatility may be endogenous. For example, governments may choose to stabilize the exchange rates with important trade partners. Thus, there may be a negative correlation between the two even though exchange rate volatility does not depress trade. Following an idea in Frankel and Wei (1994), I use the volatility of two countries' relative money supply as an instrument for the volatility of their exchange rate. As a matter of logic, this can be a good instrumental variable. Under the monetary theory of exchange rate determination, exchange rate movement is directly determined by the relative money supply. So the volatility of exchange rate should be related to the volatility of the

Table 6
Endogenous switching regressions (indirect hedging via the U.S. dollar)

	1975	1980	1985	1990	Panel I	Panel II	Panel III
Volatility 1	1.02	– 0.78	0.35	3.62*	2.58*	2.55*	2.55*
θ_1	(0.77)	(2.12)	(0.61)	(0.43)	(0.31)	(0.31	(0.31)
Volatility 2	– 4.45*	– 27.24*	– 26.41*	– 13.58*	– 30.06*	– 30.05*	– 29.35*
θ_2	(0.64)	(2.42)	(15.36)	(2.44)	(2.21)	(2.13)	(2.16)
Volatility 3						– 19.9	– 21.0
θ_3						(13.4)	(13.0)
GNP	0.77*	0.84*	0.59*	0.83*	0.74*	0.74*	0.74*
β_1	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
GNP/capita	0.24*	0.24*	0.08*	0.11*	0.20*	0.19*	0.19*
β_2	(0.02)	(0.02)	(0.03)	(0.02)	(0.01)	(0.01)	(0.01)
Distance	– 0.69*	– 0.63*	– 0.47*	– 0.62*	– 0.57*	– 0.57*	– 0.53*
β_3	(0.04)	(0.04)	(0.05)	(0.04)	(0.02)	(0.02)	(0.02)
Adjacency	0.61*	0.80*	0.79*	0.91*	0.76*	0.75*	0.73*
β_4	(0.11)	(0.17)	(0.21)	(0.17)	(0.09)	(0.09)	(0.09)
Common links	0.54*	0.61*	0.65*	0.59*	0.60*	0.59*	0.58*
β_5	(0.07)	(0.09)	(0.11)	(0.08)	(0.05)	(0.05)	(0.05)
Trade blocs							0.34*
β_6							(0.08)
Critical virtual trade Γ	7.05*	8.98*	10.01*	9.53*	10.51*	10.24*	10.26*
	(0.17)	(0.44)	(2.05)	(0.39)	(0.32)	(0.30)	(0.31)
σ_u	3.21*	3.45*	2.87*	1.42*	2.83*	2.75*	2.81*
	(0.39)	(0.80)	(1.11)	(0.54)	(0.31)	(0.31)	(0.31)
σ_e	1.22*	1.35*	1.37*	1.15*	1.24*	1.24*	1.24*
	(0.03)	(0.04)	(0.03)	(0.03)	(0.01)	(0.02)	(0.02)
ρ	0.99*	0.86*	0.91*	0.84*	0.79*	0.79*	0.78*
	(0.004)	(0.04)	(0.05)	(0.06)	(0.03)	(0.03)	(0.03)
Constant	Yes	Yes	Yes	Yes			
Year dummies					Yes	Yes	Yes
Region dummies					Yes	Yes	Yes
# obs.	1230	1457	1101	1307	5542	5542	5542
Log likelihood	– 2026.6	– 2337.5	– 2095.7	– 2200.3	– 8717.7	– 8716.5	– 8706.6

Note: (1) Standard errors are reported in parentheses.
* Statistically significantly different from zero at 5% level by a t-test.

relative money supply of the two countries in question. On the other hand, a country's monetary policy is less likely to have been manipulated to influence goods trade (unless it goes through the exchange rate).

Table 7 reports the results of regressing the exchange rate volatility on the volatility of relative money supply. We observe that the latter is clearly positively correlated with the former. The fitted values from regression reported in the last column in Table 6 are then used as an instrument in our empirical tests of the hedging hypothesis.

Table 8 replicates some key endogenous switching regressions (the last two columns in Tables 4 and 5). The results are very similar to the early ones. In particular, we find that for country pairs whose trade potentials and the size of exchange rate volatility exceed the threshold value, the volatility elasticity is numerically large (though smaller than the earlier results), negative and statistically significant. In contrast, for small country pairs, the volatility elasticity is positive. Again, the evidence does not favor the hypothesis that hedging instruments have helped to diminish the effect of exchange rate volatility on trade.

The volatility measure used in the paper is computed for the real exchange rate. Since hedging contracts are almost always designed to deal with nominal risk, one may wonder whether the results would change if nominal exchange rate volatility is used in the regressions. We may note that, since nominal exchange rate is generally much more volatile than goods prices, the volatility of the real exchange rate and that of the nominal rate should be very similar. In any case, I have redone the regressions using the nominal volatility and obtained broadly very similar results (not reported to save space).

Before concluding the battle, we should be aware of several limitations of the statistical tests. In particular, hedging instruments are also used for financial

Table 7
Volatility of relative money supply as an instrument for real exchange rate volatility

	1975	1980	1985	1990	Panel
Dependent variable: Real exchange rate volatility					
Volatility of money supply	1.304*	0.775*	2.477*	1.032*	1.509*
	(0.050)	(0.053)	(0.028)	(0.028)	(0.018)
Constant	Yes	Yes	Yes	Yes	
Year dummy					Yes
Region dummy					Yes
# obs.	1326	1378	1378	1596	5678
adj. R^2	0.34	0.14	0.85	0.46	0.57
Log likelihood	2664.1	3075.9	2294.5	2180.2	8794.6

Note: (1) Standard errors are reported in parentheses.
* Statistically significantly different from zero at 5% level.

Table 8
Using volatility of relative money supply as an instrument for real exchange rate volatility (panel regression)

	Direct hedging		Indirect hedging	
	1	2	3	4
Dependent variable = ln(Bilateral Trade)				
Volatility 1	3.51*	3.30*	3.35*	3.30*
θ_1	(0.42)	(0.51)	(0.41)	(0.40)
Volatility 2	− 12.07*	− 14.34#	− 16.47*	− 17.49*
θ_2	(2.88)	(8.46)	(4.06)	(3.62)
Volatility 3		− 5.18		− 16.41
θ_3		(8.19)		(13.69)
GNP	0.75*	0.75*	0.75*	0.75*
β_1	(0.01)	(0.01)	(0.01)	(0.01)
GNP/capita	0.18*	0.18*	0.18*	0.18*
β_2	(0.01)	(0.01)	(0.01)	(0.01)
Distance	− 0.64*	− 0.64*	− 0.63*	− 0.63*
β_3	(0.03)	(0.03)	(0.03)	(0.03)
Adjacency	0.65*	0.65*	0.64*	0.62*
β_4	(0.10)	(0.10)	(0.10)	(0.10)
Common links	0.51*	0.51*	0.48*	0.47*
β_5	(0.05)	(0.05)	(0.05)	(0.05)
Critical virtual trade Γ	9.16*	9.07*	9.38*	9.30*
	(0.55)	(0.81)	(0.41)	(0.35)
σ_u	3.08*	3.27*	1.47*	1.42*
	(0.37)	(0.48)	(0.350)	(0.33)
σ_e	1.21*	1.22*	1.20*	1.20*
	(0.02)	(0.02)	(0.01)	(0.01)
ρ	0.86*	0.85*	0.84*	0.84*
	(0.03)	(0.04)	(0.05)	(0.05)
Year dummies	Yes	Yes	Yes	Yes
Region dummies	Yes	Yes	Yes	Yes
# obs.	4484	4484	4484	4484
Log likelihood	− 7035.78	− 7035.40	− 7032.92	− 7032.28

Note: (1) Standard errors are in parentheses.
* Statistically significantly different from zero at 5% level by a *t*-test.

transactions (for speculative, if not hedging purposes). In principle, the development or the non-development of hedging markets is determined by 'virtual' financial transactions as well as virtual goods trade. Unfortunately, no data is available on a cross-section of *bilateral* financial flows. So we have maintain the assumption that the sizes of gross financial flows and of goods flows are positively correlated. It is useful to note, however, that in order to interpret the results in this paper as consistent with the hedging hypothesis, the *gross* (as opposed to net) bilateral financial flows have to be negatively correlated with the *gross* goods flows, which is not very likely.

5. Concluding remarks

Hedging instruments are commonly proposed as an explanation for the small effect of exchange rate volatility on goods trade. This paper makes a close examination of this explanation. Under the hedging hypothesis, country pairs with access to hedging instruments should suffer less from exchange rate volatility than those without hedging instruments. The difficulty with any empirical examination is a lack of good indicators that separate country pairs from one regime to the other.

Whether or not a hedging market can be developed depends on the potential demand for the instrument. Guided by this argument, the paper develops an endogenous switching regression specification. Data on more than 1000 country pairs are examined. In contrast to the hedging hypothesis, I find that country pairs with relatively small potential trade do not suffer a negative effect of volatility on trade. On the other hand, country pairs with large potential trade do exhibit a negative effect of volatility. These results cast considerable doubt on the hedging hypothesis. Various extensions and robustness checks do not overturn this basic finding.

It has been argued that many of the so-called hedging instruments are often used for speculative as opposed to true hedging purpose by currency traders (see Wei and Kim (1997) for a recent study.) The evidence in this paper is consistent with this argument. It should be noted that the statistical results in this paper are not 'constructive' in the following sense: while they help to eliminate the hedging hypothesis as an explanation of the small effect of exchange rate volatility on trade, they do not establish any clear alternative. Such will be a useful future research project.

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Appendix A. The likelihood functions for the switching regressions

A.1. Direct hedging

$$f(e|u<c)=\frac{\int_{-\infty}^{c} f(e,u)\,\mathrm{d}u}{\Pr(u<c)}=\frac{\int_{-\infty}^{c} f(e)f(u|e)\,\mathrm{d}u}{\Pr(u<c)}=\frac{f(e)\int_{-\infty}^{c} f(u|e)\,\mathrm{d}u}{\Pr(u<c)}.$$

Since (U, e) are jointly normal with correlation coefficient ρ, the distribution of U conditional on the realization of e is also normal. To be specific,

$$U|e \sim \text{Normal}\,[\rho(\sigma_u/\sigma_e)e, \sigma^2(1-\rho^2)].$$

Therefore,

$$f(e|u<c)=\frac{\dfrac{1}{\sigma_e}\phi\left(\dfrac{e}{\sigma_e}\right)\Phi\left(\dfrac{c/\sigma_u-\rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)}{\Phi\left(\dfrac{c}{\sigma_u}\right)}.$$

$f(e|u>c)$ can be worked out in a similar way.

A.2. Indirect hedging

$$f(e|u_1<c_1,u_2<c_2,u_3<c_3)=\frac{\int_{-\infty}^{c_1}\int_{-\infty}^{c_2}\int_{-\infty}^{c_3} f(e,u_1,u_2,u_3)\,\mathrm{d}u_1\,\mathrm{d}u_2\,\mathrm{d}u_3}{\Pr(u_1<c_1,u_2<c_2,u_3<c_3)}.$$

$$\text{Numerator}=\int_{-\infty}^{c_2}\int_{-\infty}^{c_3}\left[\int_{-\infty}^{c_1} f(e,u_2,u_3)f(u_1|e,u_2,u_3)\,\mathrm{d}u_1\right]\mathrm{d}u_2\,\mathrm{d}u_3.$$

Making use of the assumption that u_1 is independent of u_2 and u_3, we can rewrite the numerator as

$$\text{Numerator}=\left[\int_{-\infty}^{c_1} f(u_1|e)\,\mathrm{d}u_1\right]\int_{-\infty}^{c_2}\int_{-\infty}^{c_3} f(e,u_2,u_3)\,\mathrm{d}u_2\,\mathrm{d}u_3.$$

Proceeding in the same way, we obtain that

$$\text{Numerator}=\left[\int_{-\infty}^{c_1} f(u_1|e)\,\mathrm{d}u_1\right]\left[\int_{-\infty}^{c_2} f(u_2|e)\,\mathrm{d}u_2\right]$$

$$\times\left[\int_{-\infty}^{c_3} f(u_3|e)\,\mathrm{d}u_3\right]f(e).$$

Since $u_k \mid e$, $k = 1, 2$ and 3, is normally distributed,

$$f(e \mid I = 111) = f(e \mid u_1 < c_1, u_2 < c_2, u_3 < c_3) =$$

$$\frac{\Phi\left(\dfrac{c_1/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\Phi\left(\dfrac{c_2/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\Phi\left(\dfrac{c_3/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\dfrac{1}{\sigma_e}\phi\left(\dfrac{e}{\sigma_e}\right)}{\Phi\left(\dfrac{c_1}{\sigma_u}\right)\Phi\left(\dfrac{c_2}{\sigma_u}\right)\Phi\left(\dfrac{c_3}{\sigma_u}\right)}.$$

Other conditional densities can be derived in a similar way. I will just state the results here.

$$f(e \mid I = 121) = f(e \mid u_1 < c_1, u_2 \geq c_2, u_3 < c_3) =$$

$$\frac{\Phi\left(\dfrac{c_1/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\left[1 - \Phi\left(\dfrac{c_2/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\right]\Phi\left(\dfrac{c_3/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\dfrac{1}{\sigma_e}\phi\left(\dfrac{e}{\sigma_e}\right)}{\Phi\left(\dfrac{c_1}{\sigma_u}\right)\left[1 - \Phi\left(\dfrac{c_2}{\sigma_u}\right)\right]\Phi\left(\dfrac{c_3}{\sigma_u}\right)},$$

$$f(e \mid I = 112) = f(e \mid u_1 < c_1, u_2 < c_2, u_3 \geq c_3) =$$

$$\frac{\Phi\left(\dfrac{c_1/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\Phi\left(\dfrac{c_2/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\left[1 - \Phi\left(\dfrac{c_3/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\right]\dfrac{1}{\sigma_e}\phi\left(\dfrac{e}{\sigma_e}\right)}{\Phi\left(\dfrac{c_1}{\sigma_u}\right)\Phi\left(\dfrac{c_2}{\sigma_u}\right)\left[1 - \Phi\left(\dfrac{c_3}{\sigma_u}\right)\right]},$$

$$f(e \mid I = 122) = f(e \mid u_1 < c_1, u_2 \geq c_2, u_3 \geq c_3) =$$

$$\frac{\Phi\left(\dfrac{c_1/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{-\rho^2}}\right)\left[1 - \Phi\left(\dfrac{c_2/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\right]\left[1 - \Phi\left(\dfrac{c_3/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\right]\dfrac{1}{\sigma_e}\phi\left(\dfrac{e}{\sigma_e}\right)}{\Phi\left(\dfrac{c_1}{\sigma_u}\right)\left[1 - \Phi\left(\dfrac{c_2}{\sigma_u}\right)\right]\left[1 - \Phi\left(\dfrac{c_3}{\sigma_u}\right)\right]},$$

Finally,

$$f(e \mid I = 2) = f(e \mid u_1 \geq c_1) = \frac{\left[1 - \Phi\left(\dfrac{c_1/\sigma_u - \rho\,(e/\sigma_e)}{\sqrt{1-\rho^2}}\right)\right]\dfrac{1}{\sigma_e}\phi\left(\dfrac{e}{\sigma_e}\right)}{1 - \Phi\left(\dfrac{c_1}{\sigma_u}\right)}.$$

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